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ABSTRACT

It is a truism of research on social stratification that the effects of socioeconomic or family background on educational attainment lead to biases in the simple regression of occupational status (or other putative outcomes of schooling) on educational attainment. Using a structural equation model of sibling resemblance in educational attainment and occupational status, Hauser and Mossel have found minimal evidence of family bias in the effects of postsecondary schooling on occupational status in a sample of Wisconsin brothers. In order to resolve this seemingly anomalous finding, the present analysis compares the Hauser-Mossel findings with those in larger samples of sibling pairs of the same and of mixed sex in the Wisconsin Longitudinal Study and with pairs of brothers in Olneck's Kalamazoo study. In the course of the analysis, some methodological problems in cross-population comparisons of structural equation models are solved. The comparative analysis shows that family bias in the effects of schooling on occupational status may be much less than is commonly believed and that very large samples may be needed to measure it reliably. Moreover, the analysis suggests that estimates of family bias are very sensitive to the specification of response variability in schooling. (Author)

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Some Cross-Population Comparisons of Family Bias in the Effects of Schooling on Occupational Status

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Some Cross-Population Comparisons of Family Bias in the Effects of Schooling on Occupational Status

Abstract

It is a truism of research on social stratification that the effects of socioeconomic or family background on educational attainment lead to biases in the simple regression of occupational status (or other putative outcomes of schooling) on educational attainment. Using a structural equation model of sibling resemblance in educational attainment and occupational status, Hauser and Mossel have found minimal evidence of family bias in the effects of post-secondary schooling on occupational status in a sample of Wisconsin brothers. In order to resolve this seemingly anomalous finding, the present analysis compares the Hauser-Mossel findings with those in larger samples of sibling pairs of the same and of mixed sex in the Wisconsin Longitudinal Study and with pairs of brothers in Olneck's Kalamazoo study. In the course of the analysis, some methodological problems in cross-population comparisons of structural equation models are solved. The comparative analysis shows that family bias in the effects of schooling on occupational status may be much less than is commonly believed and that very large samples may be needed to measure it reliably. Moreover, the analysis suggests that estimates of family bias are very sensitive to the specification of response variability in schooling.

SOME CROSS-POPULATION COMPARISONS OF FAMILY BIAS
IN THE EFFECTS OF SCHOOLING ON OCCUPATIONAL STATUS

1.0 INTRODUCTION

Hauser and Mossel (1982) found minimal family bias in the effects of schooling on occupational status in a sample of 518 brothers drawn from the Wisconsin Longitudinal Study. Moreover, the regression of occupational status on years of schooling was insensitive to corrections for response variability in reports of schooling. In each sibling pair, one brother was a primary respondent who had graduated from a Wisconsin high school in 1957 and was about 36 years old at the survey date. The other brother need not have graduated from high school, but nearly all brothers (93 percent) had done so. The main difference between the populations of primary respondents and brothers is that the latter ranged from 20 to 50 years of age in 1975. Both members of each pair must have participated in a telephone survey, which was conducted in 1975 for primary respondents and in 1977 for their siblings, and both members must have held jobs in April of 1970 and at the survey dates. Most primary respondents were working in 1970, but many younger siblings were not in the labor force at that time. Thus, the siblings in this subsample are older than one might think from the nominal age restrictions. Finally, two or three measurements of educational attainment or occupational status must have been ascertained for each member of the pair.

The purpose of this analysis is to ask whether the finding of minimal family bias is at all representative of samples of siblings drawn from the Wisconsin Longitudinal Study or elsewhere. For example, as noted by Hauser and Mossel (1982:2-3), there is some evidence that bias in schooling coefficients is less in the case of post-secondary schooling than in the case of primary and secondary education. Further analysis of the Wisconsin data can shed no light on that issue, but it seems useful to establish the generality of the finding about bias within other portions of the Wisconsin sample before seeking to compare it with other, more heterogenous samples. Thus, I first obtain new estimates of omitted-variable bias in some larger and less restrictively defined samples of Wisconsin siblings. These include pairs of sisters and brother-sister pairs, as well as pairs of brothers. I also attempt to evaluate the possibility that response variability in schooling may account for observations of heterogeneity in the within- and between-family regressions of occupational status on schooling in these samples.

Second, I compare findings for the Wisconsin brother pairs with those in the Kalamazoo sample, collected by Olneck (1976, 1977). The Kalamazoo brothers are, of course, far less heterogeneous in geographic origin than those in the Wisconsin sample, but at the same time they are more heterogenous in age and in educational attainment. I first carry out unconditional tests of family bias in the Kalamazoo sample. Then, I evaluate the sensitivity of these tests to Olneck's (1976:186) estimates of response variability in schooling. Finally, I construct tests of family bias in the schooling coefficient that pool data from the Kalamazoo and Wisconsin samples. Because of the differences in the

Kalamazoo and Wisconsin study designs and populations, I carry out two different types of interpopulation comparisons. In the first set, I compare the full set of family and school effects. This comparison is excessively broad in scope, since interpopulation differences in study design, measurement, and population composition could lead to significant differences in some coefficients, even if family biases were the same in the two populations. In the second set of comparisons, I recast my model of sibling resemblance to generate a single parameter that describes family bias. Then, I test for interpopulation differences in this parameter, without having to condition on equality between populations in any other parameters of the model.

This analysis is deliberately limited in scope. First, I have not attempted to exhaust relevant bodies of data, but only to establish whether the data analyzed by Hauser and Mossel are truly unusual in the absence of family bias. Second, the analysis focuses on family bias in the effects of years of schooling on occupational status. The present analysis does not attempt to generalize to other outcomes of the stratification process, like earnings. Further, in controlling shared family effects on siblings, the analysis does not account for several other sources of omitted variable bias in the schooling coefficient. These include mental ability, school performance, social support, and aspiration, to the extent these factors vary independently within families and affect both schooling and occupational status.

2.0 A SIMPLE MODEL OF SIBLING RESEMBLANCE

Figure 1 shows a simple path model of sibling resemblance in educational attainment and occupational status. X(R) and X(S) are measures of the educational attainment of primary respondent and sibling, respectively, while Y(R) and Y(S) are the corresponding measures of occupational status. X(R) and X(S) both load on a common family education factor, K(2), while each also loads on a unique, within-family component of education, K(1) or K(3):

$$X(R) = K(2) + K(1) \tag{1}$$

and

$$X(S) = L(1)K(2) + K(3), \tag{2}$$

where $COV[K(i),K(j)] = 0$ for $i \neq j$. Similar equations define between- and within-family components of occupational status:

$$Y(R) = N(2) + N(1) \tag{3}$$

and

$$Y(S) = L(2)N(2) + N(3), \tag{4}$$

where $COV[N(i),N(j)] = 0$ for $i \neq j$. Only one of L(1) and L(2) is identified in the absence of other causes or effects of K(1) or N(2), and for the present analysis we assume L(1) = L(2) = 1 throughout. One might think this specification questionable, for example, in comparisons of the effects of parental characteristics on sons relative to daughters, but there is evidence to support it (Hauser 1983).

Last, the model specifies regressions of occupational status on educational attaiment for primary respondents, families, and siblings, respectively:

$$N(1) = G(11)K(1) + Z(1), \tag{5}$$

$$N(2) = G(22)K(2) + Z(2), \tag{6}$$

and

$$N(3) = G(33)K(3) + Z(3), \tag{7}$$

where $COV[K(i),Z(i)] = COV[K(i),Z(j)] = COV[Z(i),Z(j)] = 0$ for $i \neq j$.

3.0 ESTIMATES OF THE MODEL IN ONE POPULATION

3.1 Wisconsin Sibling Samples

Table 1 gives estimates of several versions of the model of Figure 1 in three Wisconsin samples; LISREL IV was used to obtain maximum likelihood estimates (Joreskog and Sorbom 1978). Panel A pertains to 518 pairs of Wisconsin brothers, previously analyzed by Hauser and Mossel (1982). Appendix Tables A and B describe the variables in this sample and give their means, standard deviations, and intercorrelations. Panel B pertains to a less restricted sample of 1623 pairs of Wisconsin brothers and Panel C to 598 pairs of Wisconsin sisters. These data have previously been analyzed by Hauser, Sewell, and Clarridge (1982); Appendix Table C gives the data for these two samples. Throughout the present analyses, the standard deviations of all variables have been rescaled to range from 1 to 10; this simplifies both estimation and the presentation of findings.

In each of the Wisconsin samples, educational attainments and occupational statuses of both siblings were reported by primary respondents in the 1975 survey (Clarridge, Sheehy, and Hauser 1978; Sewell and Hauser 1980). In Panels B and C, the only restrictions on the samples are that the sibling was between 20 and 65 years old and employed at the survey date. Panel A refers to a randomly selected subsample of the cases in Panel B where the brother was interviewed in 1977, and - as noted above - both the brother and the primary respondent held jobs in 1970. Thus, relative to the sample of Panel B, the selection of brothers into Panel A reflects a random subsampling process, survey nonresponse, and changes in population definition that were selective for smaller age differences between brothers.

Line A1 of Table 1 gives unrestricted estimates of the model of Figure 1 in the sample of 518 pairs of Wisconsin brothers; these same results were reported in less detail by Hauser and Mossel (1982:Table 5). Because of the differing selection criteria for primary respondents and brothers, the model of Line 1 yields distinct parameter estimates for each member of the pair. The model of Line A2 imposes the restriction, G(11) = G(33), which says that the two within-family regressions have equal slopes. The data do not violate this restriction; the likelihood-ratio test statistic, L^2, increases only by 0.42 with 1 degree of freedom (df). The model of Line A3 adds the restriction of homogeneity in the within- and between-family regressions, G(11) = G(22) = G(33); this says that there is no family bias in the effect of schooling on occupational status. As in the case of the two within-family regressions, this restriction yields a negligible decrement in fit, L^2 = 0.13 on 1 df. Line A4 equates the variances in the disturbances of occupational status for the two brothers, and Line A5 equates the variances in within-family components of occupational status. Although the study design implies a potential violation of the latter restriction, the data also fit both of these restrictions. Thus, the final model in Line A5 specifies complete symmetry between primary respondents and their brothers in the structure of the data.

Even from Line A1 of Table 1, it is evident that there is substantial homogeneity in the within- and between-family regressions in this subsample. Indeed, the unrestricted between-family regression is less steep than the within-brother regression; in the model of Line A2, the pooled within-family regression has a steeper slope than the between-family regression. Not only does this unexpected finding appear in the analysis of Panel A, but similar findings appear also in analyses of other indicators of educational attainment and occupational status within the same sample (Hauser and Mossel 1982:Table 5). This has aroused justifiable suspicions about the generality of such findings in the Wisconsin data and other bodies of sibling data.

Line B1 of Table 1 gives estimates of the model of Figure 1 for the less restricted sample of 1623 Wisconsin brothers. In this sample, the unrestricted between-family regression is apparently larger than either within-family regression. The model of Line B2 pools the two within-family slope estimates; here, there is nominally significant heterogeneity between brothers ($L^2 = 7.84$ with 1 df). Line B3 adds the restriction that all three regressions are homogeneous; the decrement in fit is nominally significant ($L^2 = 5.24$ with 1 df, $p = 0.022$). While the between-family regression is about 20 percent larger than the pooled within-family regression (Line B2), the bias is barely statistically significant in a sample that is three times larger than that of Hauser and Mossel.

Following the analysis of Panel A, the last two lines of Panel B impose additional symmetry restrictions on the parameters of the model. Here, the restriction on the disturbance variances is satisfied, but that on the within-family variances in schooling is not satisified ($L^2 = 3.96$ with 1 df).

These findings of heterogeneity and bias in the education slopes are sensitive to the specification of response variability. For example, in the sample of 518 male-male pairs, Hauser and Mossel estimated the error variances in the primary respondent's reports of his own and of his brother's schooling to be 0.496 and 0.255, respectively. When models B1 to B3 are re-estimated with these error variances introduced to correct for attenuation in schooling (without taking account of their sampling variability), the evidence of heterogeneity in slopes between brothers and between within- and between-family slopes disappears. In the model corresponding to Line B1, the test statistic is $L^2 = 0.16$ on 1 df, and the slope estimates are 0.488, 0.599, and 0.571 for respondent, family, and brother, respectively. Note that neither the fit of this model nor the between-family slope is affected by the correction for attenuation. In the model corresponding to Line B2, the test statistic increases only by $L^2 = 3.43$ on 1 df, which is not statistically significant; that is, there is no longer significant heterogeneity between the slopes for primary respondent and brother. In this model, the pooled within-family slope estimate is 0.539, and the between-family slope estimate is 0.595, so the nominal family bias is 10.4 percent. However, when the within- and between-family slope estimates are pooled, the test statistic increases only by 1.36 with 1 degree of freedom, so the family bias is not statistically significant. The final pooled slope estimate is 0.562.

This is a crude adjustment for response variability because the estimated error variances do not pertain to the full sample on which the slopes were estimated and because the findings just presented do not take account of sampling error in the estimates of response variability.

At the same time, I think that there is no substantial reason to question the borrowing of parameters in this instance, and the findings on heterogeneity and bias could only be weakened by a proper treatment of response variability.

Panel C reports an analysis of observed measurements of education and occupational status for 598 pairs of sisters in the Wisconsin sample; there are fewer pairs of sisters than brothers because of the lower labor force participation of women. As shown in Line C1, the unrestricted parameter estimates for primary respondents and their sisters are nearly symmetric, excepting the greater heterogeneity of sister's schooling. The two within-family regressions are nearly the same (compare Lines C1 and C2), and the estimated between-family slope is 22.7 percent larger than the pooled, within-family estimate. As in Panel A, however, the heterogeneity of slopes between persons and families is not statistically significant. As shown in Line C3, the additional restriction on the slopes increases the test statistic only by $L^2 = 1.64$ with 1 df. The additional symmetry restriction on disturbance variances is satisfied within the sample of Wisconsin sisters (Line C4), but the within-family variance of schooling is significantly larger among sisters than among primary respondents (Line C5).

I also re-estimated the models of Lines C1 to C3 using the same estimates of error variance in schooling that I used in correcting the results in Lines B1 to B3. I shall not report these results in detail, since they are based upon error variance estimates for male-male pairs, and even the uncorrected estimates show no family bias in the female-female pairs. However, as one might expect, these estimates show even less evidence of family bias than do the uncorrected estimates for female-female pairs.

These analyses provide weak evidence of hetergeneity in within- and between-family regressions of current occupational status on schooling in the Wisconsin sample. Among brother pairs, without taking account of response variability, the between-family regression is estimated to be 21.4 percent steeper than the (pooled) within-family regression, yet this bias is barely statistically significant in a sample nearly 5 times larger that analyzed by Hauser and Mossel. Among nearly 600 pairs of sisters, the estimated bias is 25 percent, but it is not statistically significant at even the 0.05 level. When crude estimates of response variability in schooling are used to adjust the within-family slopes, there is no significant evidence of family bias in the schooling coefficients among male-male or among female-female sibling pairs. While the Hauser-Mossel sample shows even less evidence of family bias than the two larger samples of Table 1, the evidence in those two samples is also weak. In a later section, I pool the data across these and other Wisconsin samples in an attempt to obtain a more reliable estimate of the family bias.

3.2 Kalamazoo Brothers

Table 2 gives maximum likelihood estimates of some models of sibling resemblance in educational attainment and occupational status in Olneck's (1976, 1977) data for 346 pairs of brothers from Kalamazoo, Michigan. These brothers were selected from the rolls of sixth graders in Kalamazoo public schools for the years 1928 to 1950 and were followed up in 1973. Because of sample attrition and nonresponse, the Kalamazoo data pertain to roughly one quarter of the men originally selected by Olneck; they include "only men who themselves completed an interview and who could be paired with at least one brother who also completed an interview" (Olneck 1977:127). These data are reproduced in Appendix Table D and Table E. While Olneck analyzed a number of socioeconomic outcomes, the reanalysis in Table 1 is restricted to current (1973) occupational status (Panels A and C) and to occupational status in the early career (Panels B and D). As in the Wisconsin data, educational attainment is coded in years of schooling, and occupational status is scaled on the Duncan SEI.

One important difference in the design of the Kalamazoo and Wisconsin studies is that there is no intrinsic ordering of brothers within families in the Kalamazoo study. Consequently, there is an intrinsic symmetry in the parameters of the two within-family regressions. I compensated for this within the LISREL program by placing equality constraints on equivalent parameters and reducing the nominal degrees of freedom of each model by the number of redundant moments in the variance-covariance matrix. For example, in the analysis of educational attainment and current occupation, there are nominally 10 variances and covariances, but four of these are redundant: one variance in educational attainment, one variance in occupational status, one within-brother covariance of educational attainment and occupational status, and one cross-brother covariance of educational attainment and occupational status.

Line A1 of Table 2 reports the parameters of the symmetric model of current occupational status in the Kalamazoo sample. Because of the intrinsic symmetry of the data, this baseline model fits exactly. The estimated ratio of between- to within-family slopes is 1.456, which is a good deal larger than that in any of the Wisconsin samples. At the same time, when the within- and between-family slopes are pooled (Line A2), the decrement in fit is barely statistically significant ($L^2 = 4.58$ with 1 df, $p = 0.03$). Moreover, in the case of early occupational status there is even less evidence of family bias. As shown in Line B1, the between-family slope is nominally 21.3 percent steeper than the within-family slope, but the difference between the two slopes is not statistically significant ($L^2 = 2.52$ with 1 df, $p = 0.11$).

Corrections for response variability in educational attainment and in occupational status are introduced in the models of Panel C and D. No parameters for response variability are identified in the model of Figure 1 as it stands, but Olneck (1976:186) reports correlations between true and observed values of educational attainment (0.964), status of early occupation (0.896), and status of current occupation (0.917). Using these correlations, I estimated components of error variance in the observed variables and introduced these variance components into the model. For example, educational attainment has a standard deviation of 2.73, and I estimated its error variance as

$$2.73^2 (1 - 0.964^2) = .5269.$$

The error variance components were introduced here as constants, and their sampling variablility has been ignored. Consequently, as in my nominally corrected estimates for the two Wisconsin samples, the sampling variability of the corrected estimates has been underestimated to an unknown degree.

As shown in Panels C and D, there is less evidence of family bias in the Kalamazoo data when they have been corrected for response variability. Again, the between-family slope estimates are indifferent to the corrections for response variability until they are pooled with the within-family slopes; the correction for attenuation only increases the within-family slopes and, thereby, the pooled slopes. In the case of current occupational status, the corrected ratio of between- to within-family slopes is 1.23, but the difference in slopes is not statistically significant. In status of early occupation, the corrected slopes are virtually identical. Thus, the evidence of family bias in slopes of occupational status on schooling does not appear to be any more convincing in the Kalamazoo data than in the Wisconsin data.

4.0 POOLING ESTIMATES ACROSS SAMPLES

Lack of statistical power may be one explanation of the unreliability of our estimates of bias. Now that statistical methods of testing for bias are available, our first real finding may be that the major sibling samples are not large enough to measure it. Of course, another possibility is that the family bias is not very large or important. One way to help choose between these explanations is to pool estimates of bias across existing samples. That is, estimate a model of sibling resemblance simultaneously in samples from two or more nominally comparable populations and pool the estimates of family bias. Here, I first pool estimates of family bias across four sex-types of sibling pairs within the Wisconsin sample. Second, I pool estimates of bias for brother pairs in the Wisconsin and Kalamazoo samples.

4.1 Four Wisconsin Sibling Samples

Table 3 gives estimates of models of sibling resemblance in educational attainment and occupational status in 1975 for independent samples of Wisconsin siblings in each of the four possible sex combinations. The sex of the primary respondent is listed first in each combination, so male-female and female-male pairs are distinct. The same-sex pairs have already been introduced in the analysis of Table 1; there are 797 male-female pairs and 1020 female-male pairs. The data are reports by the primary respondent about her or his own educational attainment and occupation in 1975 and those of a randomly selected sibling. The main reason for variation across sex combinations in the number of observations is the lower labor force participation of women; in addition, there appears to be a tendency for male respondents to underreport the labor force activity of their sisters. In all, there are data for 4038 distinct sibling pairs.

There are two reasons to suppose that the present estimates of family bias may be too large. First, as confirmed in the preceding analyses, a correction for response variability in schooling will increase the within-family slopes. Second, to the extent that measured family background characteristics affect both the common factor in schooling and that in occupational status, introduction of such background variables may reduce the between-family slope. The latter correction permits a decomposition of total family bias into portions attributable to measured and unmeasured family characteristics, but does not reduce the global estimate of bias.

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The first panel of Table 3 reports estimates of the basic model of Figure 1 in the four Wisconsin samples with no cross-group constraints on parameters. This model imposes one overidentifying restriction within each group by virtue of the normalization of equal loadings on the family factor for the primary respondent and the sibling (Hauser and Mossel 1982), but the departure of the data from these restrictions is negligible (L^2 = 2.58 with 4 df). Estimates of this model for the same-sex pairs have already been presented in Table 1. There is prima facie evidence of family bias in the fact all of the between-family slope estimates are larger than any of the within-family estimates. The largest between-family slope, 0.664 in male-male pairs, is nearly three times larger than the smallest within-family slope, 0.242 among female respondents in mixed pairs. There is also a good deal of variability across groups in the within-family slope estimates (and in other parameter estimates); the largest within-family slope, 0.539 among male siblings in brother pairs, is more than two times larger than the smallest within-family slope. There appears to be less variability across samples in the between-family slopes, which range only from 0.545 in female-female pairs to 0.664 in male-female pairs.

In order to simplify and strengthen tests of family bias, I have pooled a number of parameter estimates across the four sibling subgroups. In one model (not shown in Table 3), equality constraints were placed on all parameters that pertained to persons of the same sex and response status. For example, this reduced the number of within-family slope estimates from 8 to 4, and the same reduction was made in the number of parameters for within-family variances in schooling and in the number of parameters for within-family disturbances in occupational status. This set of 12 restrictions yielded a test statistic (relative to the model of Panel 1) of $L^2 = 63.39$. One might well think of the violations of these restrictions as of little substantive importance, and it is thus instructive to find that the sample is large enough for such trivial restrictions to yield a highly significant test statistic.

Panel 2 of Table 3 shows the parameters of a model in which cross-sample equality constraints on parameters of the between-family regressions have been added to those on sex and response-status specific parameters of the within-family models. Relative to the model with only the latter restrictions, the fit statistic increases by 15.21 with 9 df, which is not statistically significant at even the 0.05 level. Again, one might think of these restrictions on between-family parameters as substantively trivial, for all of the Wisconsin sibling pairs have been drawn from the same population of families. To put the matter in a slightly weaker fashion, many of the families from which one pair-type was drawn at random might also have contributed some other pair-type to the sample.

As shown in Panel 2, these restrictions have reduced the apparent differences in within-family slopes and in within-family disturbance variances for each sex between primary respondents and their siblings. For this reason, I imposed additional equality restrictions on the sex-specific within-family slopes and within-family disturbance variances, but not on the within-family variances in educational attainment. That is, the model specifies equal slopes and disturbance variances for brothers, regardless of response status, and it specifies equal slopes and disturbance variances for sisters, regardless of response status; it does permit variances of schooling to vary by sex and response status. Panel 3 of Table 3 shows these constrained estimates. Relative to the model of Panel 2, the fit deteriorates by 9.78 with 4 df, which is nominally significant with $p = 0.04$. Given the observed similarity of the unconstrained parameters in Panel 2 and the more significant test statistics associated with several trivial restrictions, I have conditioned the remainder of the analysis on these pooled estimates. The advantage in doing so is that there are now just two estimates of family bias, described by a ratio of 0.603/0.462 = 1.305 in the case of men and by a ratio of 0.603/0.359 = 1.680 in the case of women.

Conditional on the model of Panel 3 in Table 3, it is straightforward to test the statistical significance of these estimates of family bias. The specification of equal between-family slopes and male, within-family slopes yields a test statistic of $L^2 = 19.32$ with 1 df, and the specification of equal between-family slopes and female, between-family slopes yields a test statistic of $L^2 = 34.78$ with 1 df. Both contrasts are highly significant. Thus, by pooling more than 4000 Wisconsin sibling pairs and ignoring response variability in schooling, it has been possible to produce statistically significant test statistics for family bias in occupation-schooling regressions.

In one test of the sensitivity of these findings, I obtained an estimate of response variability in schooling from the measurement error model of Hauser and Mossel (1982:19-25) for Wisconsin brother pairs. I pooled the estimates of response variances in the primary respondent's reports of his own and of his brother's schooling; the two estimates were not significantly different ($L^2 = 3.30$ with 1 df). The pooled estimate is 0.379 with a standard error of 0.023. Since the within-family variances in schooling differ substantially by sex and response status, so also do the corrections in slopes implied by this single estimate of response variability in schooling. For example, the estimate implies a reliability of 0.908 in the schooling of female respondents and a reliability of 0.945 in the schooling of male respondents.

When this estimate of response variance is introduced into the model of Panel 3 in Table 3, the fit improves slightly relative to the model without the correction (L^2 = 89.10 with 29 df). The estimate of family bias declines to 20.6 percent among men and 41.1 percent among women. Conditional on this specification of response variability, the restriction of equality in the between-family and male, within-family slopes yields a test statistic of L^2 = 9.73 with 1 df, and the restriction of equality in the between-family and female, within-family slopes yields a test statistic of L^2 = 14.90 with 1 df. Both contrasts remain statistically significant, but they are less convincing than before. Interestingly, since the unrestricted within-family slope estimate for males (0.501) falls between that for females (0.428) and the between-family slope estimate (0.604), there is little deterioration in the fit of the model if one adds the specification of completely homogeneous slopes to that of homogeneity in the female, within-family slope and the between-family slope (L^2 = 0.70 with 1 df). In the model of complete homogeneity, the pooled slope estimate is 0.530 with a standard error of 0.009.

I think it reasonable to suppose that the preceding adjustment for response variability is minimal, that is, that the correct adjustment might be larger and the evidence of family bias correspondingly weaker. For example, among 1452 Wisconsin men who were employed in-state in 1964, Massagli and Hauser (1981:Table 9) estimated the error variance in son's schooling as 0.608. As noted earlier, Olneck's (1976:186) estimate of the reliability (0.929) of schooling reports among Kalamazoo brothers implies an error variance of 0.527; my own analysis of his repeated measurements of educational attainment and occupational status yields a larger point estimate, 0.862. Also, Bielby and Hauser (1977:262) report estimates of error standard deviations in years of schooling of 1.08 and 1.03 among 813 nonblack male respondents to the March 1973 Current Population Survey who also completed the Income Supplement Reinterview, and they report estimates of 0.97, 1.78, and 0.60 for three measures of schooling among 556 nonblack male respondents to the March 1973 CPS who also completed the Occupational Changes in a Generation Supplement and Reinterview. All but one of these various estimates is substantially larger than that employed in the preceding calculations.

Rather than basing further calculations on another of these point estimates, I turn the problem around and ask how large an estimate of response variability in schooling is required to account for the observed heterogeneity in within- and between-family regressions. That is, conditional on homogeneity of within- and between-family regressions, random error variance components in schooling are identified. Distinct estimates for primary respondent and sibling are identified within each of the subsamples, but for the present purpose I have pooled these estimates within sex and response-status combinations. Thus, in the least restrictive model of this form, I estimate response variance components in the schooling of male respondents, 1.358 (0.241), male siblings, 1.000 (0.248), female respondents 0.954 (0.158), and female siblings 0.973 (0.196); standard errors are given in parentheses. The fit of this model is slightly better than that of Panel 3 in Table 3 (L^2 = 85.23 with 27 df). Under this model, the pooled estimate of the slope is 0.606 (0.023). Further, the differences among the 4 estimates of the response variance in schooling are not statistically significant. When I pool the estimates across response status within sex, the test statistic increases by L^2 = 2.72 with 2 df; the pooled estimates are 1.181 (.217) for men and 0.965 (0.141) for women. When I pool these two estimates, the test statistic increases by 1.32 with 1 df; the pooled estimate of response variance in schooling is 0.972 (0.143). In this model the pooled estimate of the regression of occupational status on schooling is 0.593 (0.019). The overall fit of the model, L^2 = 89.27 with 30 df, is virtually the same as that of Panel 3 in Table 3.

If one accepts this estimate of response variability - which falls well within the range of other published estimates - there is no need to specify either family biases or sex differentials in occupational returns to schooling. I think it most interesting that a single estimate of response variability in schooling may account for the two, quite distinct estimates of family bias that were obtained for males and females. The reason, mentioned briefly above, is that the observed variance in schooling, as well as the within-family regression of occupational status on schooling, is less among females than males. For example, the pooled estimate of the total variance in schooling among female respondents is 4.140, so the implied reliability of schooling is 1 - 0.972/4.140 = 0.765. Among male respondents, the pooled estimate of variance in schooling is 6.894, and the implied reliability is 1 - 0.972/6.894 = 0.859. The specification that reports of schooling by or about men and women are equally accurate - in the sense that the scatter of observations about true values is the same - leads to far larger adjustments in estimated slopes for women than for men. In fact, the implied correction in female, within-family slopes is larger than needed to eliminate the family bias; recall that the initial estimates of response variance in schooling were lower for females than males. If I estimate a single error variance component in schooling by specifying no family bias among men, but not necessarily among women, the resulting estimate of the female slope, 0.618, is actually larger than that of the pooled male and between-family slope, 0.595; however, those two slope estimates are not significantly different from each other.

Aside from the issues surrounding heterogeneity in slopes, there is another interesting sidelight to the possibility that the error variance in schooling may be close to 1.0. It substantially affects our estimates of the relative heterogeneity of families with respect to schooling outcomes. For example, in the model of Panel 3 in Table 3, the between-family variance in schooling, 2.254, is less than half the within-family variance among male primary respondents or siblings. If the error variance in schooling is 0.972, then the between-family variance component is more than half the true within-family variance. The effect is much larger among females. In the case of female respondents, the observed within-family variance in schooling is 83 percent as large as the between-family variance, but the corrected within-family variance is only 40 percent as large as the between-family variance. In the case of female siblings, the observed within-family variance in schooling is 27 percent larger than the between-family variance, but the corrected within family variance is 84 percent as large as the between-family variance. Not only do these corrections alter our estimates of the relative variability of schooling outcomes within and between families, but they also illustrate what is to me, at least, a non-obvious fact, that there is inverse variation between two important family effects on achievement; the greater the homogeneity of educational outcomes within families, the less is the family bias in the effects of schooling on occupational status.

It is not clear quite how seriously one ought to take these results. I think they should carry more weight, say, than Griliches' (1979:S53-S54) "back of the envelope" efforts to show that response variability may account for observed family biases in earnings functions for men. At the same time, I make no claim that they provide the last word on the subject, even in the Wisconsin data. The data actually include multiple measurements of the educational attainments of respondents and siblings in subsamples of all four sex-pair types, not merely those analyzed by Hauser and Mossel (1982). Further analyses of those data should permit a more convincing specification of the relative importance of family bias and response variability in the observed heterogeneity of within and between-family regressions of occupational status on schooling. The importance of pursuing these analyses is indicated by the fact that they may also help to explain sex differentials in the effects of schooling.

4.2 Kalamazoo And Wisconsin Brothers

By comparing the regressions of occupational status on schooling in the 518 Wisconsin brother pairs of Hauser and Mossel (1982) with those in the 346 Kalamazoo brother pairs of Olneck (1976, 1977), I hope to increase the generality, as well as the statistical power of my estimates of family bias. In both these samples, all of the observations are based on self-reports by each member of the pair. The samples differ because of the local character of the Kalamazoo sample, the greater range of ages it includes, and the absence of any explicit truncation of the schooling distribution; moreover, the Kalamazoo data, but not the Wisconsin data, are intrinsically symmetric.

Panel A of Table 4 reports the estimates of a model, like that of Figure 1, in which no cross-population constraints have been imposed on the Wisconsin and Kalamazoo data. However, to simplify the analysis, I have imposed constraints of complete symmetry on the Wisconsin data, which account for the imperfect fit of the model. Obviously, there are a number of differences in the estimated parameters of the model in the Wisconsin and Kalamazoo data. For example, the within-family slope estimate is nearly 75 percent larger in the Wisconsin than in the Kalamazoo data, and the between-family variance in educational attainment is twice as large in Kalamazoo as in Wisconsin. Panel B gives a set of pooled estimates for Kalamazoo and Wisconsin, estimated under the assumption that all of the parameters are the same in both populations. This model does not fit ($L^2 = 71.74$ with 10 df), nor is there any reason for it to fit, given the several differences between the two samples that I have just enumerated. However, conditional on this specification of complete homogeneity between the two populations, the additional restriction of equal within- and between-family regressions does not lead to a significant deterioration in fit. As shown in Panel C, the test statistic increases only by 0.79 with 1 df.

In an effort to condition the pooled test for bias on a more acceptable model, I specified a model in which the within- and between-family slopes, but no other coefficients were equal in the Wisconsin and Kalamazoo samples. These estimates are shown in Panel D of Table 4. Again, the data are inconsistent with the cross-sample equality constraints; the contrast of the model of Panel D with that of Panel A yields a test statistic of $L^2 = 24.66$ with 2 df. As shown in Panel E, the additional constraint of homogeneity in the within- and between-family regressions does not lead to any further deterioration in fit.

In order to avoid pooling heterogeneous coefficients as a condition of the test of family bias, I recast the basic model of sibling resemblance to include a single parameter for the ratio of within-family to between-family slopes. That is, the model is written to include a single parameter whose value is the reciprocal of the measure of family bias that I have used descriptively throughout the preceding text. Using this model, it is possible to pool estimates of family bias across samples and to test the significance of the pooled estimate of bias without conditioning on cross-sample equality of any other coefficients in the model.

Figure 2 shows a path diagram of the revised model. In the revised model, there is a single slope, G, for the regressions within and between families. However, rather than specifying unit slopes in the regressions of individual educational attainment on its family components, the model specifies a free parameter, L, which takes on the same value for the primary respondent and sibling. If G and L are each free parameters of the model, G is the within-family slope of the model of Figure 1, and L is the ratio of within-family to between-family slopes in that model. In the setup of Figure 1, to test the hypothesis that there is no family bias, it is necessary to place an equality constraint on the within- and between-family slopes and to compare the fit under that constraint with the fit of a model excluding that constraint. In the setup of Figure 2, the same hypothesis can be tested simply by observing whether a confidence interval about L includes 1.0.

Panel A of Table 5 reports the estimates of this model with no cross-sample constraints in the Kalamazoo and Wisconsin data. The model is equivalent to that of Panel A in Table 4. Notice that the reparameterization of the model rescales the between-family variance in schooling and that the bias parameter, L, is in each sample the ratio of the previously estimated within-family and between-family slopes. A 95 percent confidence interval about the bias parameter includes 1.0 in the Wisconsin sample, but not in the Kalamazoo sample.

Panel B of Table 5 gives estimates of the model under the constraint that the bias parameter, but no other parameter of the model, is the same in the two populations. This constraint increases the test statistic by $L^2 = 2.92$ with 1 df, so we are unable to reject the hypothesis that family bias is the same in the populations from which the Kalamazoo and Wisconsin samples were drawn. In Panel C of Table 5, the pooled bias parameter is fixed at unity, which says there is no family bias in either population. Under this model, the test statistic

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increases by $L^2 = 1.68$ with 1 df relative to that under the model of equal bias; we are unable to reject the hypothesis that there is no family bias. The same hypothesis could have been tested simply by noting that 1.0 is included in a 95 percent confidence interval about the estimate of bias in Panel B.

In short, without having to specify homogeneity in any parameters of sibling resemblance between Kalamazoo and Wisconsin brothers, I have found that there is no substantial evidence that family bias differs in those two populations. Moreover, even without any adjustment for response variability in schooling in either population, I have found that the pooled estimate of family bias is negligible.

5.0 CONCLUSION

In Hauser and Mossel's sample of 518 Wisconsin brother pairs there was negligible evidence of family bias in the regressions of occupational status on schooling, whether or not the data were corrected for response variability in reports of completed schooling. In light of our theoretical expectation that family bias would be substantial, and of the appearance of such bias in other samples, the present analysis has investigated family biases in other Wisconsin sibling samples and in Olneck's sample of Kalamazoo brothers.

The Hauser-Mossel sample does appear to be unusual in the fact that there is so little prima facie evidence of family bias in the schooling coefficient. The present analysis suggests that family biases are subject to a great deal of sampling variability, so it should come as no surprise when they fail to appear in a sample. For example, in an

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analysis of current occupational status (Table 5) in the Hauser-Mossel sample a two-sigma confidence interval about the ratio of within- to between-family slopes ranges from 1.347 to 0.703, and in the Kalamazoo sample a two-sigma confidence interval about this ratio ranges from 0.938 to 0.434.

Even in samples where there is prima facie evidence of family bias, it may not be statistically reliable. For example, the bias is significant just beyond the 0.05 level in the case of current occupational status in the Kalamazoo sample, but in the same sample it is not statistically significant in the case of occupational status of the first job. In a much larger sample of 1623 Wisconsin brother pairs, the observed ratio of between-family to within-family slopes is 1.21, but that bias is barely significant at the 0.05 level. In a sample of 598 Wisconsin sister pairs, a slightly larger observed bias, 1.25, is not statistically significant. In a comparative analysis of current occupational status among Kalamazoo and Wisconsin men, there was no substantial evidence of family bias. Only by pooling sex-specific bias estimates across samples of more than 4000 Wisconsin sibling pairs was it possible to obtain statistically reliable evidence of family bias.

Moreover, where heterogeneity did appear in within- and between-family regressions of occupational status on schooling, it may well have been an artifact of response variability in schooling. In the sample of 1623 Wisconsin men, a modest correction for response variability eliminated the family bias. In the Kalamazoo sample, the family bias disappeared when Olneck's estimate of the reliability of schooling was introduced. In the pooled sample of 4038 Wisconsin sibling pairs, the family bias was substantially reduced by introducing a minimal estimate of response variability in schooling. In the same pooled sample, the specification of absolutely no family bias implied an estimate of response variability in schooling that was consistent with other published estimates.

My reading of the present evidence is that family biases in the effects of schooling on occupational status are a good deal weaker and more variable than most investigators, including myself, have previously thought. It will be interesting to see whether sampling variability also accounts for the uneven evidence of family bias in earnings functions. Recall Griliches' (1979:S54) complaint that "something else [besides response variability] must be going on." Another matter worth pursuing elsewhere is the evidence of omitted variable bias in schooling regressions with explicit measures of family background. Should we believe the ubiquitous evidence that the schooling coefficient declines under controls for socioeconomic background? It is time to take that question seriously.

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FIGURE 1. A structural equation model of sibling resemblance in educational attainment and occupational status

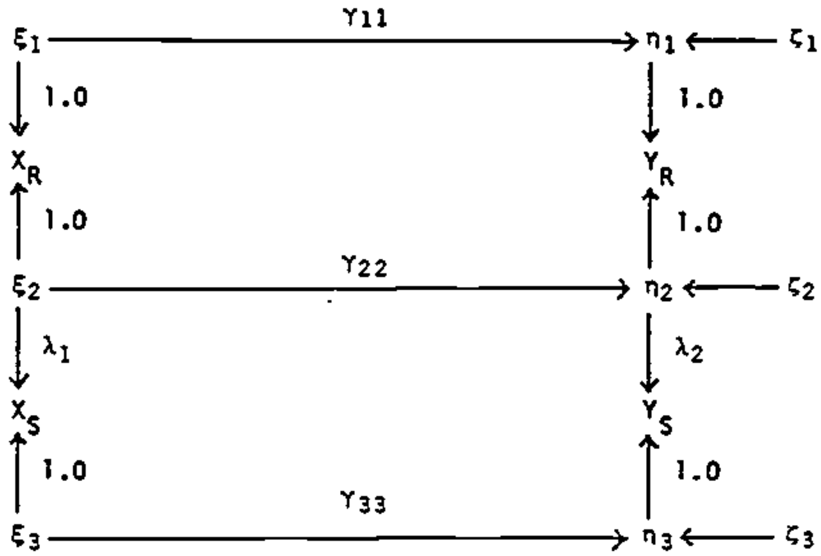

FIGURE 2. A structural equation model of sibling resemblance in educational attainment and occupational status with a parameter for family bias

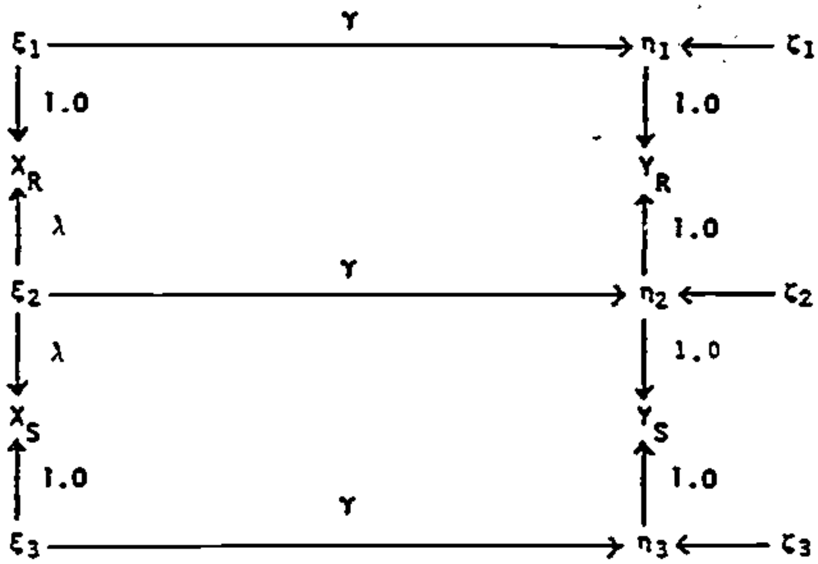

Table 1. Maximum likelihood estimates of models of sibling resemblance in educational attainment and occupational status: Wisconsin samples.

Sample and model	Slopes			Var. in ed. att.			Var. in occ. status			Measures of fit		
	Resp. γ_{11}	Fam. γ_{22}	Sib γ_{33}	Resp. ϕ_1	Fam. ϕ_2	Sib ϕ_3	Resp. ψ_1	Fam. ψ_2	Sib ψ_3	L^2	df	p
A. Wisconsin sample: 518 male-male pairs (from Hauser-Mossel 1982)												
1. $\lambda_1 = \lambda_2 = 1$	0.636	0.638	0.697	2.432	1.955	2.988	3.325	0.811	3.248	0.73	1	0.39
	(0.074)	(0.073)	(0.062)	(0.240)	(0.222)	(0.263)	(0.272)	(0.184)	(0.269)			
2. Add $\gamma_{11} = \gamma_{33}$	0.672	0.635	0.672	2.432	1.950	2.993	3.323	0.815	3.248	1.15	2	0.56
	(0.048)	(0.072)	(0.048)	(0.240)	(0.222)	(0.263)	(0.272)	(0.184)	(0.268)			
3. Add $\gamma_{11} = \gamma_{22} = \gamma_{33}$	0.658	0.658	0.658	2.437	1.950	2.988	3.320	0.816	3.251	1.28	3	0.73
	(0.030)	(0.030)	(0.030)	(0.240)	(0.222)	(0.263)	(0.272)	(0.184)	(0.269)			
4. Add $\psi_1 = \psi_3$	0.658	0.658	0.658	2.437	1.950	2.988	3.286	0.816	3.286	1.32	4	0.86
	(0.030)	(0.030)	(0.030)	(0.240)	(0.222)	(0.263)	(0.204)	(0.184)	(0.204)			
5. Add $\phi_1 = \phi_3$	0.658	0.658	0.658	2.713	1.950	2.713	3.286	0.816	3.286	3.51	5	0.62
	(0.030)	(0.030)	(0.030)	(0.169)	(0.222)	(0.169)	(0.204)	(0.184)	(0.204)			
B. Wisconsin sample: 1623 male-male pairs (from Hauser-Sewell-Clarridge 1982)												
1. $\lambda_1 = \lambda_2 = 1$	0.425	0.599	0.539	3.849	2.969	4.482	3.165	0.511	3.151	0.18	1	0.67
	(0.032)	(0.031)	(0.028)	(0.210)	(0.191)	(0.224)	(0.145)	(0.093)	(0.143)			
2. Add $\gamma_{11} = \gamma_{33}$	0.490	0.595	0.490	3.886	2.961	4.451	3.194	0.516	3.144	8.02	2	0.02
	(0.022)	(0.031)	(0.022)	(0.210)	(0.192)	(0.224)	(0.144)	(0.094)	(0.143)			
3. Add $\gamma_{11} = \gamma_{22} = \gamma_{33}$	0.530	0.530	0.530	3.848	2.964	4.489	3.207	0.528	3.146	13.26	3	0.00
	(0.013)	(0.013)	(0.013)	(0.210)	(0.192)	(0.225)	(0.145)	(0.093)	(0.143)			
4. Add $\psi_1 = \psi_3$	0.529	0.529	0.529	3.848	2.964	4.489	3.177	0.528	3.177	13.37	4	0.01
	(0.013)	(0.013)	(0.013)	(0.210)	(0.192)	(0.225)	(0.112)	(0.093)	(0.112)			
5. Add $\phi_1 = \phi_3$	0.529	0.529	0.529	4.169	2.964	4.169	3.177	0.528	3.177	17.33	5	0.00
	(0.013)	(0.013)	(0.013)	(0.146)	(0.192)	(0.146)	(0.112)	(0.093)	(0.112)			

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Table 1. Continued.

Sample and model	Slopes Resp. γ_{11}	Slopes Fam. γ_{22}	Slopes Sib γ_{33}	Var. in ed. att. Resp. ϕ_1	Var. in ed. att. Fam. ϕ_2	Var. in ed. att. Sib ϕ_3	Var. in occ. status Resp. ψ_1	Var. in occ. status Fam. ψ_2	Var. in occ. status Sib ψ_3	Measures of fit L^2	df	p
C. Wisconsin sample: 598 female-female pairs (from Hauser-Sewell-Clarridge 1982)												
1. $\lambda_1 = \lambda_2 = 1$	0.435 (0.076)	0.545 (0.054)	0.437 (0.052)	1.807 (0.185)	1.964 (0.190)	2.734 (0.220)	3.044 (0.217)	0.137 (0.127)	2.832 (0.107)	1.44	1	0.23
2. Add $\gamma_{11} = \gamma_{33}$	0.436 (0.045)	0.545 (0.053)	0.436 (0.045)	1.807 (0.184)	1.964 (0.190)	2.733 (0.219)	3.044 (0.217)	0.137 (0.127)	2.832 (0.207)	1.44	2	0.49
3. Add $\gamma_{11} = \gamma_{22} = \gamma_{33}$	0.485 (0.025)	0.485 (0.025)	0.485 (0.025)	1.798 (0.184)	1.966 (0.190)	2.743 (0.220)	3.057 (0.217)	0.143 (0.126)	2.833 (0.207)	3.08	3	0.38
4. Add $\psi_1 = \psi_3$	0.486 (0.025)	0.486 (0.025)	0.486 (0.025)	1.798 (0.184)	1.966 (0.190)	2.743 (0.220)	2.945 (0.170)	0.143 (0.126)	2.945 (0.170)	3.87	4	0.42
5. Add $\phi_1 = \phi_3$	0.486 (0.025)	0.486 (0.025)	0.486 (0.025)	2.270 (0.131)	1.966 (0.191)	2.270 (0.131)	2.946 (0.170)	0.142 (0.127)	2.946 (0.170)	13.42	5	0.02

Note: Parenthetic entries are standard errors. Wisconsin data are based on reports by the primary respondent in the 1975 survey.

Table 2. Maximum likelihood estimates of models of sibling resemblance in educational attainment and occupational status: Kalamazoo brothers (N=346).

Dependent variable and model	Slopes: Brother $\gamma_{11} = \gamma_{33}$	Slopes: Family γ_{22}	Var. in ed. att.: Brother $\phi_1 = \phi_3$	Var. in ed. att.: Family ϕ_2	Var. in occ. status: Brother $\psi_1 = \psi_3$	Var. in occ. status: Family ψ_2	Measures of fit: L^2	Measures of fit: df	Measures of fit: p
A. Current occupational status: Uncorrected estimates									
1. Symmetry	0.401 (0.052)	0.584 (0.049)	3.361 (0.256)	4.092 (0.458)	3.170 (0.241)	0.262 (0.191)	0	0	1.00
2. $\gamma_{11} = \gamma_{22} = \gamma_{33}$	0.498 (0.027)	0.498 (0.027)	3.361 (0.256)	4.092 (0.458)	3.201 (0.244)	0.292 (0.189)	4.58	1	0.03
B. Early occupational status: Uncorrected estimates									
1. Symmetry	0.559 (0.045)	0.678 (0.044)	3.361 (0.256)	4.092 (0.458)	2.384 (0.181)	0.351 (0.151)	0	0	1.00
2. $\gamma_{11} = \gamma_{22} = \gamma_{33}$	0.620 (0.024)	0.620 (0.024)	3.361 (0.256)	4.091 (0.458)	2.397 (0.182)	0.366 (0.150)	2.52	1	0.11
C. Current occupational status: Corrected for response variability									
1. Symmetry	0.475 (0.062)	0.584 (0.049)	2.834 (0.256)	4.092 (0.458)	2.215 (0.243)	0.262 (0.191)	0	0	1.00
2. $\gamma_{11} = \gamma_{22} = \gamma_{33}$	0.539 (0.029)	0.539 (0.029)	2.802 (0.252)	4.122 (0.458)	2.196 (0.242)	0.290 (0.188)	1.32	1	0.25
D. Early occupational status: Corrected for response variability									
1. Symmetry	0.662 (0.055)	0.678 (0.044)	2.834 (0.256)	4.092 (0.458)	1.072 (0.184)	0.351 (0.151)	0	0	1.00
2. $\gamma_{11} = \gamma_{22} = \gamma_{33}$	0.671 (0.026)	0.671 (0.026)	2.827 (0.252)	4.099 (0.456)	1.067 (0.182)	0.354 (0.149)	0.03	1	0.86

Note: Parenthetic entries are standard errors; these are underestimated in Panels B and C because of failure to take account of sampling variability in response variances. Data are from Olneck (1977) with error variance estimates based on Olneck (1976:186).

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Table 3. Maximum likelihood estimates of models of sibling resemblance in educational attainment and occupational status: Wisconsin sample subgroups.

Model and Subgroup	Slopes			Var. in ed. att.			Var. in occ. status		
	Resp. γ_{11}	Fam. γ_{22}	Sib γ_{33}	Resp. ϕ_1	Fam. ϕ_2	Sib ϕ_3	Resp. ψ_1	Fam. ψ_2	Sib ψ_3
1. No cross-subgroup or within subgroup constraints: L^2 = 2.58 with 4 df, p = 0.63									
Male-male	0.425	0.599	0.539	3.849	2.969	4.482	3.165	0.511	3.151
(N=1623)	(0.032)	(0.031)	(0.028)	(0.210)	(0.191)	(0.224)	(0.145)	(0.093)	(0.143)
Male-female	0.408	0.664	0.347	5.771	1.895	3.172	3.907	0.022	2.690
(N=797)	(0.035)	(0.065)	(0.049)	(0.356)	(0.231)	(0.263)	(0.229)	(0.128)	(0.182)
Female-female	0.435	0.545	0.437	1.807	1.964	2.734	3.044	0.137	2.832
(N=598)	(0.076)	(0.054)	(0.052)	(0.185)	(0.190)	(0.220)	(0.217)	(0.127)	(0.207)
Female-male	0.242	0.634	0.424	1.985	2.111	5.429	2.865	0.138	3.904
(N=1020)	(0.064)	(0.049)	(0.031)	(0.183)	(0.186)	(0.288)	(0.175)	(0.118)	(0.205)
2. Equal parameters for same sex and response-status: L^2 = 81.18 with 25 df, p = 0.00									
Male-male	0.427	0.608	0.488	4.683	2.262	5.011	3.518	0.239	3.565
	(0.022)	(0.023)	(0.020)	(0.181)	(0.098)	(0.177)	(0.118)	(0.056)	(0.115)
Male-female	0.427	0.608	0.390	4.683	2.262	2.903	3.518	0.239	2.637
	(0.022)	(0.023)	(0.033)	(0.181)	(0.098)	(0.161)	(0.118)	(0.056)	(0.119)
Female-female	0.304	0.608	0.390	1.850	2.262	2.903	2.863	0.239	2.637
	(0.047)	(0.023)	(0.033)	(0.122)	(0.098)	(0.161)	(0.122)	(0.056)	(0.119)
Female-male	0.304	0.608	0.488	1.850	2.262	5.011	2.863	0.239	3.565
	(0.047)	(0.023)	(0.020)	(0.122)	(0.098)	(0.177)	(0.122)	(0.056)	(0.115)
3. Add equal sex-specific slopes and disturbance variances: L^2 = 90.96 with 29 df, p = 0.00									
Male-male	0.462	0.603	0.462	4.710	2.254	4.994	3.542	0.247	3.542
	(0.015)	(0.023)	(0.015)	(0.181)	(0.098)	(0.176)	(0.089)	(0.056)	(0.089)
Male-female	0.462	0.603	0.359	4.710	2.254	2.872	3.542	0.247	2.762
	(0.015)	(0.023)	(0.028)	(0.181)	(0.098)	(0.160)	(0.089)	(0.056)	(0.091)
Female-female	0.359	0.603	0.359	1.882	2.254	2.872	2.762	0.247	2.762
	(0.028)	(0.023)	(0.028)	(0.122)	(0.098)	(0.160)	(0.091)	(0.056)	(0.091)
Female-male	0.359	0.603	0.462	1.882	2.254	4.994	2.762	0.247	3.542
	(0.028)	(0.023)	(0.015)	(0.122)	(0.098)	(0.176)	(0.091)	(0.056)	(0.089)

Note: See text for explanation.

Table 4. Maximum likelihood estimates of models of sibling resemblance in educational attainment and occupational status: Pooled data for Wisconsin (N=518) and Kalamazoo (N=346) brothers.

Model	Slopes		Var. in ed. att.		Var. in occ. status	
	$\gamma_{11} = \gamma_{33}$	γ_{22}	$\psi_1 = \psi_3$	ψ_2	$\psi_1 = \psi_3$	ψ_2
A. No restrictions (except symmetry in the Wisconsin data): $L^2 = 6.63$ with 4 df, $p = 0.16$						
Kalamazoo	0.401 (0.052)	0.584 (0.049)	3.361 (0.256)	4.092 (0.458)	3.170 (0.241)	0.262 (0.191)
Wisconsin	0.682 (0.047)	0.665 (0.074)	2.854 (0.178)	1.923 (0.226)	3.289 (0.205)	0.601 (0.183)
B. Same coefficients in each population: $L^2 = 71.74$ with 10 df, $p = 0.00$						
Pooled estimates	0.558 (0.035)	0.618 (0.043)	3.057 (0.147)	2.791 (0.221)	3.301 (0.159)	0.589 (0.134)
C. Same coefficients, homogeneous regressions: $L^2 = 72.53$ with 9 df, $p = 0.00$						
Pooled estimates	0.584 (0.020)	0.584 (0.020)	3.057 (0.147)	2.791 (0.221)	3.302 (0.159)	0.592 (0.134)
D. Same slopes in each population: $L^2 = 31.29$ with 6 df, $p = 0.00$						
Kalamazoo	0.572 (0.035)	0.585 (0.041)	3.365 (0.256)	4.084 (0.457)	3.268 (0.249)	0.267 (0.191)
Wisconsin	0.572 (0.035)	0.585 (0.041)	2.850 (0.177)	1.929 (0.227)	3.323 (0.207)	0.612 (0.165)
E. Same slopes, homogeneous regressions: $L^2 = 31.33$ with 7 df, $p = 0.00$						
Kalamazoo	0.578 (0.020)	0.578 (0.020)	3.361 (0.256)	4.092 (0.455)	3.275 (0.249)	0.262 (0.191)
Wisconsin	0.578 (0.020)	0.578 (0.020)	2.854 (0.178)	1.923 (0.226)	3.319 (0.206)	0.815 (0.165)

Note: See text for explanation.

Table 5. Maximum likelihood estimates of models of sibling resemblance in educational attainment and occupational status with 1 parameter for heterogeneity in within- and between-family regressions: Pooled data for Wisconsin (N=518) and Kalamazoo (N=346) brothers.

Model	Slopes		Var. in ed. att.		Var. in occ. status	
	γ	λ	$\phi_1 = \phi_3$	ϕ_2	$\psi_1 = \psi_3$	ψ_2
A. No restrictions (except symmetry in the Wisconsin data): L^2 = 6.63 with 4 df, p = 0.16						
Kalamazoo	0.401 (0.052)	0.686 (0.126)	3.361 (0.256)	8.696 (3.274)	3.170 (0.241)	0.262 (0.191)
Wisconsin	0.682 (0.047)	1.025 (0.161)	2.854 (0.178)	1.829 (0.614)	3.289 (0.205)	0.800 (0.183)
B. Equal λ in Kalamazoo, Wisconsin: L^2 = 9.55 with 5 df, p = 0.089						
Kalamazoo	0.461 (0.039)	0.863 (0.098)	3.353 (0.255)	5.513 (1.389)	3.182 (0.242)	0.291 (0.188)
Wisconsin	0.639 (0.040)	0.863 (0.098)	2.863 (0.178)	2.561 (0.647)	3.294 (0.205)	0.792 (0.184)
C. λ = 1 in Kalamazoo, Wisconsin: L^2 = 11.23 with 6 df, p = 0.082						
Kalamazoo	0.498 (0.027)	1.000 ---	3.361 (0.256)	4.092 (0.458)	3.201 (0.244)	0.292 (0.189)
Wisconsin	0.676 (0.029)	1.000 ---	2.854 (0.178)	1.923 (0.226)	3.289 (0.205)	0.801 (0.183)

Note: See text for explanation.

Table A. Description of the variables, mnemonics, source of report, and year of measurement: Wisconsin brothers (N = 518)

Mnemonic	Description	Source	Year
1. EDEQYR	Respondent's Years of Schooling	Respondent	1975
2. EDAT64	Respondent's Years of Schooling	Parent	1964
3. XEDEQYR	Sib's Years of Schooling	Sibling	1977
4. SSBED	Sib's Years of Schooling	Respondent	1975
5. OCSXCR	Respondent's Current Occupation	Respondent	1975
6. OCSX70	Respondent's 1970 Occupation	Respondent	1975
7. XOCSXCR	Sib's Current Occupation	Sibling	1977
8. OCSSIB	Sib's 1975 Occupation	Respondent	1975
9. XOCSX70	Sib's 1970 Occupation	Sibling	1977

Note: Occupation is scaled on Duncan's Socio-Economic Index.

Table B. Product-moment correlation coefficients, means, and standard deviations: Wisconsin brothers (N = 518)

Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
1. EDEQYR	1.000								
2. EDAT64	0.906	1.000							
3. XEDEQYR	0.404	0.437	1.000						
4. SSBED	0.419	0.450	0.926	1.000					
5. OCSXCR	0.552	0.525	0.251	0.252	1.000				
6. OCSX70	0.590	0.562	0.300	0.295	0.818	1.000			
7. XOCSXCR	0.217	0.243	0.622	0.568	0.264	0.315	1.000		
8. OCSSIB	0.217	0.245	0.627	0.593	0.265	0.307	0.815	1.000	
9. XOCSX70	0.228	0.257	0.628	0.575	0.247	0.275	0.819	0.780	1.000
Mean:	13.60	13.38	13.37	13.29	4.91	4.88	4.80	4.72	4.49
St.Dev:	2.09	1.83	2.27	2.22	2.44	2.41	2.57	2.51	2.54

Note: Correlations are based on 518 pairs of brothers for whom complete data were available. For explanation of mnemonics, see Table 1. For convenience in the scaling of coefficients, values of the Duncan SEI have been divided by 10.

Table C. Correlations among educational attainments and occupational statuses of siblings by sex composition of sibling pair

		Respondent		Sibling		
		Ed	Oc	Ed'	Oc'	Std. dev.
<u>Male respondents</u>						
Respondent	Ed	—	.562	.416	.272	2.61
	Oc	.545	—	.285	.276	23.35
Sibling	Ed'	.308	.229	—	.625	2.73
	Oc'	.234	.181	.529	—	24.53
Std. dev.		2.77	23.91	2.25	19.84	797 1623
<u>Female respondents</u>						
Respondent	Ed	—	.482	.467	.248	1.94
	Oc	.455	—	.274	.180	20.41
Sibling	Ed'	.379	.263	—	.519	2.17
	Oc'	.258	.208	.545	—	20.05
Std. dev.		2.02	20.00	2.75	24.14	1020 598

Note: Entries above diagonal pertain to same-sex siblings, and those below diagonal to opposite-sex siblings. We use Ed and Oc for achievements of respondents and Ed' and Oc' for those of siblings. Base frequencies are given in the lower right hand corner of each panel. Source is Hauser, Sewell and Clarridge (1982: Table 18).

Table D. Means and Standard Deviations for Kalamazoo Brothers Complete Data Sample (N = 346 Pairs)

Variable	Mean	S.D.
1. Father's Education (FATH ED)	9.51	3.33
2. Father's Occupation (FATH OC)	38.33	22.52
3. Siblings (N SIBS)	3.72	2.53
4. Education (ED1,ED2)	13.20	2.73
5. Early Occupation (FJOB1,FJOB2)	39.51	23.80
6. Current Occupation (OCC1,OCC2)	49.91	23.17
7. Respondent's Income or Earnings (EARN1,EARN2)	16746	7634
8. Natural Logarithm of Income or Earnings (LN$1,LN$2)	9.62	.45

NOTE: The Kalamazoo measure is respondent's expected 1973 annual earnings. Source is Olneck (1976:37).

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Table E. Correlations among background and achievement variables: Kalamazoo Brothers (N = 346 Pairs)

	LN$1	EARN1	OCC1	FJOB1	ED1	IQ1
LN$1	1.000					
EARN1	0.938	1.000				
OCC1	0.409	0.482	1.000			
FJOB1	0.386	0.411	0.563	1.000		
ED1	0.407	0.431	0.591	0.716	1.000	
IQ1	0.360	0.359	0.453	0.445	0.576	1.000
AGE1	-0.083	-0.071	-0.105	-0.140	-0.184	-0.164
LN$2	0.220	0.219	0.218	0.211	0.269	0.169
EARN2	0.219	0.237	0.225	0.231	0.285	0.178
OCC2	0.218	0.225	0.309	0.321	0.378	0.300
FJOB2	0.211	0.231	0.321	0.394	0.427	0.326
ED2	0.269	0.285	0.378	0.427	0.549	0.400
IQ2	0.169	0.178	0.300	0.326	0.400	0.469
AGE2	-0.050	-0.032	-0.120	-0.142	-0.157	-0.158
FATH ED	0.160	0.171	0.215	0.350	0.400	0.261
FATH OC	0.197	0.212	0.218	0.391	0.383	0.260
N SIBS	-0.154	-0.155	-0.220	-0.256	-0.328	-0.276

	AGE1	LN$2	EARN2	OCC2	FJOB2	ED2
AGE1	1.000					
LN$2	-0.050	1.000				
EARN2	-0.032	0.938	1.000			
OCC2	-0.120	0.409	0.482	1.000		
FJOB2	-0.142	0.386	0.411	0.563	1.000	
ED2	-0.157	0.407	0.431	0.591	0.716	1.000
IQ2	-0.158	0.360	0.359	0.453	0.445	0.576
AGE2	0.587	-0.083	-0.071	-0.105	-0.140	-0.184
FATH ED	-0.182	0.160	0.171	0.215	0.350	0.400
FATH OC	-0.165	0.197	0.212	0.218	0.391	0.383
N SIBS	0.066	-0.154	-0.155	-0.220	-0.256	-0.328

	IQ2	AGE2	FATH ED	FATH OC	N SIBS
IQ2	1.000				
AGE2	-0.164	1.000			
FATH ED	0.261	-0.182	1.000		
FATH OC	0.260	-0.165	0.470	1.000	
N SIBS	-0.276	0.066	-0.250	-0.224	1.000

Note: Source is Olneck (1977:131-132).

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